Financial Contagion and the European Debt Crisis

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Abstract

Since the beginning of 2010, the Euro Area faces a severe sovereign debt crisis, now generally known as the Euro Crisis. While the Euro Crisis has its origin in Greece, other countries meanwhile experienced very similar problems. Dynamic conditional correlation models (DCC) are estimated for government yield spreads of selected Euro Area countries in order to assess if contagion is identifiable during the Euro Crisis, or if the countries' problems are instead due to global shocks or fundamental problems in the affected economies. Our findings suggest that there is contagion at work within the Euro Area. Specifically, contagion effects generated by negative rating announcements are documented. These results are crucial when it comes to choosing the correct measure and timing of policy intervention.

JEL classification: E43, E44, E63

Keywords: Contagion, Dynamic conditional correlation, Euro Crisis

1. Introduction

Since the beginning of 2010 the Euro Area faces a severe sovereign debt crisis, now generally known as the Euro Crisis. Rising government deficits and debt levels triggered rating agencies to downgrade several European countries' debt repayment probabilities, thereby creating a broader loss of confidence in financial markets. At the same time bond yields increased considerably further worsening the repayment abilities of ailing countries. The creation of the European Financial Stability Facility on 9 May 2010 and the intervention of the International Monetary Fund did neither reverse the widening of the yield spreads vis-à-vis the German government yield nor contain the crisis to Greece. While the Euro Crisis finds its origin in Greece which was the first country to be rescued with loans from other Euro area members and the IMF, Figure 1 shows that yields of many other Euro Area countries have meanwhile increased substantially as well. Our paper focuses on the question whether these increases in yields are caused by fundamental factors or instead due to contagion.

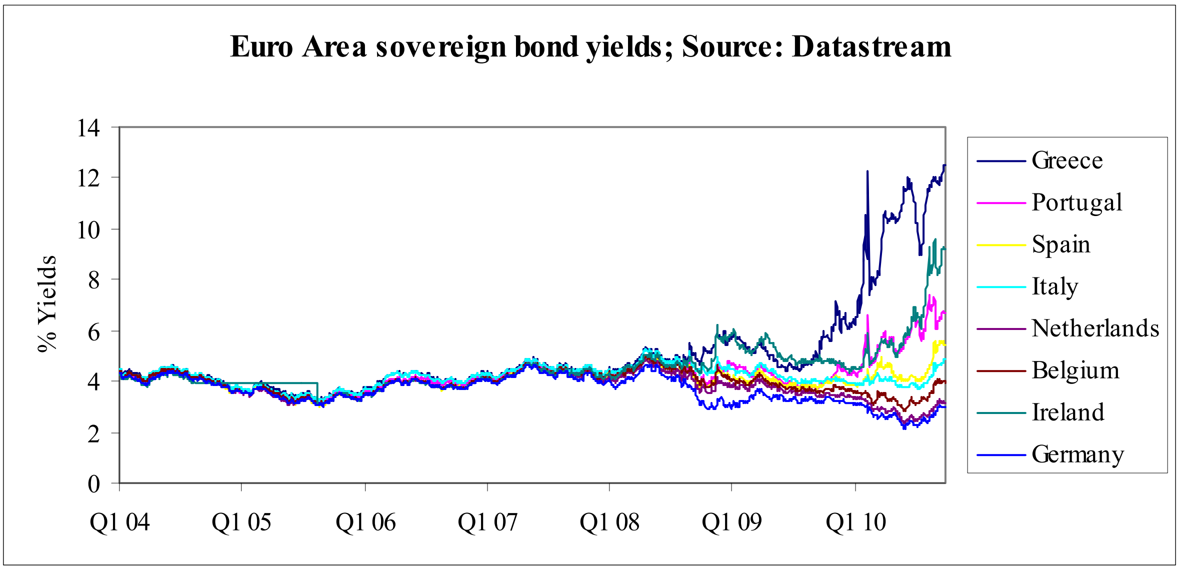

Figure 1: Bond yields of selected sovereigns of the Euro Area.

Essentially the rising yields of some countries other than Greece might result from financial markets recognizing that those countries are themselves fundamentally in severe trouble. However, it might also be the case that Greece somehow "infected" other countries by negatively influencing the markets' assessment of countries which are otherwise financially rather sound. It is this latter mechanism which we mean by contagion. The question if the current refinancing problems of some countries are disproportionate to their actual fundamental problems is therefore a question of contagion. This paper applies multivariate

GARCH-models, specifically the Dynamic Conditional Correlation (DCC) model of Engle and Sheppard (2001), to study the correlation dynamics of Euro Area bond yields. This allows us to empirically study the dynamics of the bond yield correlations and how they react to idiosyncratic shocks to the Greek economy.

Following Kaminsky et al. (2003) we define contagion as the "fast and furious" reaction of financial market prices in asset markets or countries to events that are – or seem at the time – unrelated to the fundamental environment of the reacting markets or countries. We argue below that rating downgrades for Greek government bonds provide us with reasonably unrelated events for other Euro Area government bonds. Studying the correlation of the various yields around such rating announcements should thus shed light on the contagion hypothesis.

The main findings of the paper support the view that contagious effects were at work during the summer of 2010. Specifically, our results show that five countries (Belgium, Ireland, Italy, Portugal, Spain) out of a sample of six (the Netherlands being the remaining one) were affected by contagion. Furthermore, we examine if negative rating announcements for Greece – which we argue were exogenous to the other countries – generated these contagious effects. The results are ambiguous, but there are at least strong tendencies that downgrades for Greece were translated to Ireland, Portugal and Spain, while the other sample countries remained largely unaffected from Greek rating cuts.

To study contagion we adopt a multivariuate GARCH methodology, specifically the DCC model. We focus on volatility and correlation measures for two reasons. First, stylized facts of contagion transmission mechanisms as summarized in Corsetti et al. (2010) support an analysis in that direction. All four facts, (i) spreading stock price declines across countries, (ii) increasing return volatility in crisis periods, (iii) generally higher covariance, and (iv) sometimes higher correlations in times of financial turmoil, are justifying an investigation of contagion by looking at the volatilities or correlations of financial assets.

Second, the literature on contagion also supports this approach. Several working definitions for contagion have been used in the literature. Pericoli and Sbracia (2003) provide a summary of the five most commonly used ones: Contagion can be found if (i) the probability of a crisis in one country rises sharply as a response to a crisis in another country, (ii) the increase in asset price volatility is cross-national, (iii) comovements of asset prices are not fundamentally driven, (iv) the comovements of financial assets between countries increase significantly and (v) the transmission mechanism between countries changes conditional on a crisis in one of the countries, also leading to a change in the comovement of those countries' asset prices. Definitions (ii), (iii), (iv) and (v) all analyse contagion by comparing volatilities or correlations of financial assets. In particular, we rely on working definitions (iii) and (iv) which are most in line with the Kaminsky et al. (2003) definition of "fast and furious" contagion and which can be empirically tested by means of DCC analysis.

Crucial contributions to the contagion literature also aim at that direction. Identifying periods of higher correlations between financial assets as contagious periods, King and Wadhwani (1990) were among the first authors to employ this kind of contagion analysis. Relying on a more sophisticated heteroskedasticity adjusted correlation measure Forbes and Rigobon (2002) also investigate comovements in different markets in order to distinguish between normal interdependence caused by economic relationships and contagion caused by changing investor sentiment, herding behaviour or panic. Carporale et al. (2005) further advance the contagion analysis by endogenizing the break points between normal and contagious periods thereby rendering the static analysis dynamic. Finally, Chiang et al. (2010) identify contagion and herding effects for the Asian crisis in 1997 by estimating a Dynamic Conditional Correlation (DCC) model.

The paper is organized as follows: Section 2 describes the Dynamic Conditional Correlation model (DCC) in more detail. Section 3 presents the data and the estimation specification. Our results and some cautious policy implications are provided in Section 4. Section 5 extends the analysis to study if announcements of Greek rating downgrades had any contagious effects on other countries' yields. Section 6 concludes.

2. Model

DCC models were introduced by Engle (2002) and Engle and Sheppard (2001). They belong to the class of multivariate generalized autoregressive conditional heteroskedasticity (MGARCH) models. Specifying MGARCH models is not easy, as a compromise between parsimonious but still interpretable models needs to be found. Further the models need to be constructed in a way that guarantees symmetry and positive definiteness of the estimated covariance or correlation matrices. A DCC specification fulfils all requirements from an econometric point of view, i.e. the estimation is feasible for a high amount of assets being examined and nevertheless parsimonious. The estimated correlation matrices are guaranteed to be positive definite. Additionally, the conditional variances and conditional correlations are estimated directly. As those volatility corrected correlation dynamics are of key interest in the contagion analysis, the results from DCC estimation are also nicely interpretable from an economic point of view.

In the DCC model, univariate GARCH models are first estimated for each single asset return and the standardized residuals from the models for the conditional variances are then used to calculate the conditional correlations. The detailed procedure is presented in the following section.

2.1 Model Setup

In a first step, mean zero returns r_t with covariance matrix H_t derived from some filtration (e.g. ARMA residuals) are used to estimate a GARCH specification of the conditional variances for all k assets in the analysis. In what follows a generic asset return r_t is assumed to follow a GARCH process according to equation (1):

$$h_t = \omega + \Sigma_p \alpha_p r^2_{t-p} + \Sigma_q \beta_q h_{t-q} \tag{1}$$

Here h_t represents the conditional variances, r_t the filtered mean-zero returns and ω, the α's and β's the parameters to be estimated. The filtered return process can be written according to (2).

$$r_t = h_t \varepsilon_t \qquad \text{with} \quad \varepsilon_t \sim N(0,1) \quad \text{and} \quad r_t \sim N(0, H_t) \tag{2}$$

With the estimates of the univariate GARCH equations in (1), the conditional variances h_t can be used to derive the standardized GARCH residuals ε_t from (2). Those standardized residuals are required to model the dynamic correlation structure. Specifically, the correlation dynamics are estimated according to the DCC equation (3) and the normalizition (4).

$$Q_t = (1 - \Sigma_m \gamma_m - \Sigma_n \delta_n)\,\hat{O} + \Sigma_m \gamma_m (\varepsilon_{t-m} \varepsilon'_{t-m}) + \Sigma_n \delta_n Q_{t-n} \tag{3}$$

$$R_t = Q^{*}{}_t^{-1} Q_t Q^{*}{}_t^{-1} \quad (4)$$

Q_t represents the time varying covariance matrix of the standardized residuals; Ô the unconditional covariance matrix of the standardized residuals, the γ's and the δ's the estimated parameters of the DCC equation. As in a GARCH equation, the covariance dynamics depends on past shocks and past covariances. The required lag length of the DCC equation (3) again has to be identified. The γ's represent the reaction of the comovement to innovations, i.e. to past shocks, whereas the δ's represent the decay of past comovements. The unconditional covariance matrix Ô is positive definite and the lagged shocks $\varepsilon_{t-m}\varepsilon'_{t-m}$ are positive semidefinite, consequently also Q_t as weighted average of a positive definite and a positive semidefinite matrix will be positive definite. Engle and Sheppard (2001) provide exact conditions for this result.

The normalization (4) is used to arrive at the dynamic correlation matrices R_t. $Q^{*}{}_t$ is a diagonal matrix with the square roots of the diagonal of Q_t as diagonal elements. By multiplying with the inverse, the typical element of R_t is the correlation coefficient of two assets and the diagonal of R_t consists of ones as the correlation of one asset with itself necessarily equals one.

Finally, the still unspecified time varying covariance matrix of the filtered returns r_t is derived according to (5).

$$H_t = D_t R_t D_t \quad (5)$$

Here, D_t is a diagonal matrix with the square roots of the estimated conditional variances h_t as typical element.

The non diagonal elements of R_t represent the estimated correlation coefficients between the analysed assets. Those correlation coefficients are ideal for contagion identification for the following reasons.

Firstly, correlations of assets are generally crucially dependent on the volatility of the assets as argued for example by Boyer et al. (1999), Loretan and English (2000) or Corsetti et al. (2005). As stylized facts demonstrate that volatility is higher in crisis times, this snaps through to the correlation measures. Therefore examining the correlation without controlling for the change in volatilities is problematic. The standardized residuals ε_t used in DCC equation (3) and the normalizition (4) are defined as the filtered returns r_t divided by the conditional variances h_t. Consequently, the correlations are estimated by using volatility adjusted input variables ε_t. Therefore it is guaranteed that changes in the level of correlations are not solely driven by changes in volatilities.

Secondly, the dynamic structure of the contagious effects plays an important role. One test might fail to deliver correct results, because contagious tendencies are present, however can only be recognized some lags later. Therefore, when correcting for heteroskedasticity, the dynamic structure of the variance process should be accounted for as argued by Hong (2001). The same is true for the dynamic structure of the correlation coefficients. By using a DCC model, the exact lag length of both the conditional volatility (1) and the conditional comovements (3) can be modelled.

Thirdly, no static correlation coefficients need to be used. By calculating static correlation coefficients or static test statistics for different subperiods, the test results crucially depend on the timing of those subperiods. The exact break of the transmission mechanism, i.e. the starting point of the contagion needs to be clearly identifiable, as different break points might generate different results. Consequently, by estimating models based on such exogenous assumptions, the dynamic structure of the contagious effects might get lost, as correct inference strongly depends on those assumptions regarding the data generating processes. By using the dynamic correlation structure from DCC estimation, no exogenous assumptions regarding the timing of contagious periods need to be made, but contagious periods can be determined endogenously.

2.2 Model Estimation

The DCC model can be estimated using maximum likelihood. If the input variables r_t are not multivariate normal, quasi maximum likelihood is applied instead. A two stage estimation procedure can be use to solve for the parameters maximizing the likelihood function (6).

$$L = -1/2\ \Sigma_t\ [\ k \log(2\pi) + \log\left|H_t\right| + r_t'\ H_t^{-1} r_t\ ] =$$

$$= -1/2\ \Sigma_t\ [\ k \log(2\pi) + 2 \log\left|D_t\right| + r_t'\ D_t^{-2}\ r_t - \varepsilon_t'\varepsilon_t + \log\left|R_t\right| + \varepsilon_t'\ R_t^{-1}\ \varepsilon_t\ ] \qquad (6)$$

In the first stage, the univariate GARCH equations for all k assets are estimated using the filtered return series r_t as input variables. By maximizing the volatility part (7) of the likelihood function jointly for all k assets, the GARCH parameters ω, α and β for each asset and each asset's lag order are estimated.

$$L_{GARCH} = -1/2\ \Sigma_t\ [\ k \log(2\pi) + 2 \log \left| D_t \right| + r_t'\ D_t^{-2}\ r_t\ ] \quad (7)$$

Using these parameters, the conditional variances h_t can be derived, which are required to calculate the standardized residuals ε_t according to (2). With the standardized residuals, the second stage, i.e. the DCC estimation can be conducted. Therefore, the DCC part (8) of the likelihood function is maximized conditional on the parameter estimates from the first stage. Here, the DCC parameters γ and δ for the required lag order are estimated.

$$L_{DCC} = -1/2\ \Sigma_t\ [-\ \varepsilon_t'\varepsilon_t + \log \left| R_t \right| + \varepsilon_t'\ R_t^{-1}\ \varepsilon_t\ ] \quad (8)$$

Under very general conditions, the two stage (quasi) maximum likelihood estimates are consistent and asymptotically normal as can be found in White (1994). Additionally, Bollerslev-Wooldridge consistent standard errors can be calculated according to Engle and Sheppard (2001). Consequently, consistent t-tests on the parameters of both estimation stages can be conducted.

3. Dataset and Specification

For the investigation of contagious effects during the Euro Crisis we use a sample of seven countries vis-à-vis Germany: Belgium, Greece, Ireland, Italy, Portugal, Spain and the Netherlands. Thus, our sample of countries includes both countries which witnessed refinancing problems resulting from increased government bond yields as well as countries seemingly unaffected by the crisis. Additionally, we study those countries' yields vis-à-vis the German yield because the German yield serves as our riskfree benchmark and because of the sheer size of the German economy within the Euro Area. The exclusion of such a major country in a model of contagion would lead to errors due to misspecification as shown by Dungey et al. (2003).

3.1 Dataset

A dataset consisting of daily 10-year benchmark government bid yields calculated by Thomson Reuters and provided on Datastream for a time period from 01/01/2009 until 12/31/2010 is applied in the analysis. Using daily returns for a five-day week, the sample covers a total of 522 data points. The benchmark yield represents that yield, which countries need to offer for newly emitted 10-year government bonds in order to attract investors. German benchmark yields are used as a reference point in order to compare the risk premium of the affected countries. By subtracting the German benchmark yields, parallel developments of monetary policy and parallel inflation expectations of the countries examined are removed from the benchmark yields. Consequently, we choose yield spreads in our empirical analysis as they represent the country specific risk premium of Belgium, Greece, Ireland, Italy, Portugal, Spain and the Netherlands.

As mentioned in Subsection 2.1, the input variables for a DCC model need to have an expected return of zero according to (2), therefore the yield spreads from the original dataset need to be modified by some kind of filtration. As the focus of the analysis lies on the dynamic process of the second moments of the spread series, an ARMA-filtration for the original data is used in order to provide input variables not being autocorrelated with respect to the first moment.

ADF and KPSS tests for unit root identification are conducted for each time series. As every series is integrated of order one, the subsequent analysis is done in first differences. The government bond yield time series show clear signs for a structural break in 2008. However we focus exclusively on the time span from 01/01/2009 until 12/31/2010 for which no break needs to be modelled. The ARMA lag-length selection is identified via Hannan-Rissanen model selection and the Schwarz information criterion. Models are checked for no remaining autocorrelation using Portmanteau and LM tests. Applying the procedure to the seven spread series results in a filtration of the first differences according to ARMA(1,2) for Belgium, ARMA(7,1) for Greece, MA(1) for Ireland, MA(2) for Italy, MA(4) for Portugal, AR(3) for Spain and AR(2) for the Netherlands. Remaining autocorrelation can clearly be rejected for all time series examined. Additionally, the filtered spreads nearly always exhibit signs for conditional heteroskedasticity as verified with ARCH-LM tests.[1] Consequently, a GARCH analysis is both reasonable from a theoretical point of view in order to investigate contagious effects and from an econometrically point of view.

[1] Portugal represents the only exception as conditional heteroskedasticity of the MA(4) residuals is clearly rejected by the ARCH-LM test.

3.2 Specification

In order to estimate the dynamics of the second order moments and the correlation structure, the filtered spreads are applied in a DCC model. The specification of the DCC model requires the identification of the lag order for the single univariate GARCH equations provided in (1) and for the multivariate DCC equation provided in (3).

For the specification of the GARCH order, nested versions of the models for each time series are sequentially evaluated with likelihood ratio tests. Alternatively, it would also be possible to choose the GARCH order according to the minimization of information criteria. Residuals are checked for remaining conditional heteroskedasticity using modified Portmanteau and LM tests for no remaining ARCH-effects. This methodology identifies a GARCH(1,2) process for Belgium, GARCH(1,1) process for Greece, GARCH(1,1) for Ireland, GARCH(1,1) for Italy, GARCH(4,5) for Portugal, GARCH(1,3) for Spain and GARCH(1,3) for the Netherlands. The resulting residuals are significantly tested for no remaining conditional heteroskedasticity.

The specification of the DCC equation (4) requires more attention. As explained in Subsection 2.1, the DCC equation is used to estimate the correlation dynamics from the assets of interest. These correlation dynamics are required to analyse the potential presence of contagious effects. However, as a first step, we should investigate if there are correlation dynamics at all or if the correlation remains constant as assumed in Bollerslev (1990). If that was the case, contagion could be rejected right from the beginning of the analysis and only questions of interdependence could be answered. As argued in Forbes and Rigobon (2002), high comovements of markets are only a sign of strong economic linkages between those countries. If those comovements are similarly high during crisis and non-crisis periods, one should only conclude for interdependence of those countries, as transmission mechanisms remain fairly stable. Only if the transmission and the cross-market linkages get propagated during crisis times, contagion will be at hand. Engle and Sheppard (2001) propose a test which evaluates the null of constant conditional correlation against dynamic conditional correlation. The OLS based test demonstrates good size and power properties against local alternatives. With a p-value of approximately zero[2], the null of constant conditional

[2] An extensive overview of all statistical results can be obtained from the authors upon request.

correlation can clearly be restricted at any level of significance. Thus, estimation of a dynamic correlation structure seems indeed necessary.

Secondly, it needs to be analysed if the dynamics of the comovements have a unit root and therefore an integrated dynamic conditional correlation model should be estimated instead of a regular DCC model. This can be accomplished by estimating a restricted model which imposes $1 - \delta = \gamma$ and comparing it via a likelihood ratio test to the unrestricted DCC(1,1) specification. An integrated DCC model can clearly be rejected.

Finally, in order to determine the lag length of the DCC equation sequential likelihood ratio tests can again be conducted as in the univariate GARCH specification. However, those tests can only be used as an initial analysis, as the resulting test statistics are not necessarily Chi-squared distributed, as demonstrated in Foutz and Srivastave (1977). The identified specification should be verified with Wald tests, as those are consistent due to the modified standard errors. Using this test procedure the DCC process is identified to be DCC(1,1).

4. Results

4.1 Correlation Dynamics

Using the DCC model as specified in Subsection 3.2, six correlation series are calculated for each country of the sample. Greece was the first country witnessing the problems of the sovereign debt crisis and therefore represents the origin of our contagion analysis. Specifically, we study contagion by examining the pairwise dynamic correlations of government bond yield spreads for Belgium, Ireland, Italy, Portugal, Spain and the Netherlands vis-à-vis Greece. Our sample of countries thus includes some of the Euro Area periphery and some of the core. This allows us to study whether there were any contagious effects at work, and if so whether there was a difference between periphery and core countries.

As argued in the introduction we define contagion here as the “fast and furious” reaction of government bond yields of countries other than Greece to events only related to the Greek economy. Our DCC model should then detect this by finding rising correlations immediately following a Greek event. If the DCC results indicate that contagious effects between Greece and other countries were present, then financial market participants – at least to some degree – transferred the financial problems of Greece to other countries in the Euro Area that would otherwise be fundamentally sound. Instead if no contagion effects are found, then the Greek situation is independent from the other countries' financial development.

The dynamic correlation structure between Greece and the other six countries of the sample is provided in Figure 2.[3] It shows the daily evolution of the bond spread correlation between 2009 and 2011. Both DCC and CCC models are estimated. The constant horizontal line depicts the bond spread correlation of the CCC model, i.e. under the assumption of constant correlation over the sample period. This assumption can however be rejected for the data, as already discussed in Section 3.2. The time-varying line represents the dynamic correlation structure. Greek government bond spreads always exhibit positive correlations with the other countries' spreads.

An overview of the descriptive statistics of the correlation structure is provided in Table 1. On average the correlation between Greek and Portuguese yield spreads is highest with a mean of 0.542. With an average of 0.315 the correlation between Greek and Dutch yield spreads is the lowest over the entire sample period. While significantly different from constant, all correlation dynamics are nevertheless not very erratic with standard deviations ranging from 0.067 (Greece-Netherlands) to 0.082 (Greece-Belgium).

	Minimum	Maximum	Mean	SD
Greece-Belgium	0.174	0.746	0.458	0.082
Greece-Ireland	0.212	0.745	0.495	0.074
Greece-Italy	0.237	0.728	0.498	0.074
Greece-Portugal	0.340	0.773	0.542	0.068
Greece-Spain	0.252	0.760	0.463	0.070
Greece-Netherlands	0.144	0.488	0.315	0.067

Table 1: Greek correlation dynamics: Descriptive statistics.

[3] The analysis aims at investigating if contagious effects are generated with Greece being the origin, therefore only the Greek correlations are displayed here. The results for the other countries can be obtained from the authors upon request.

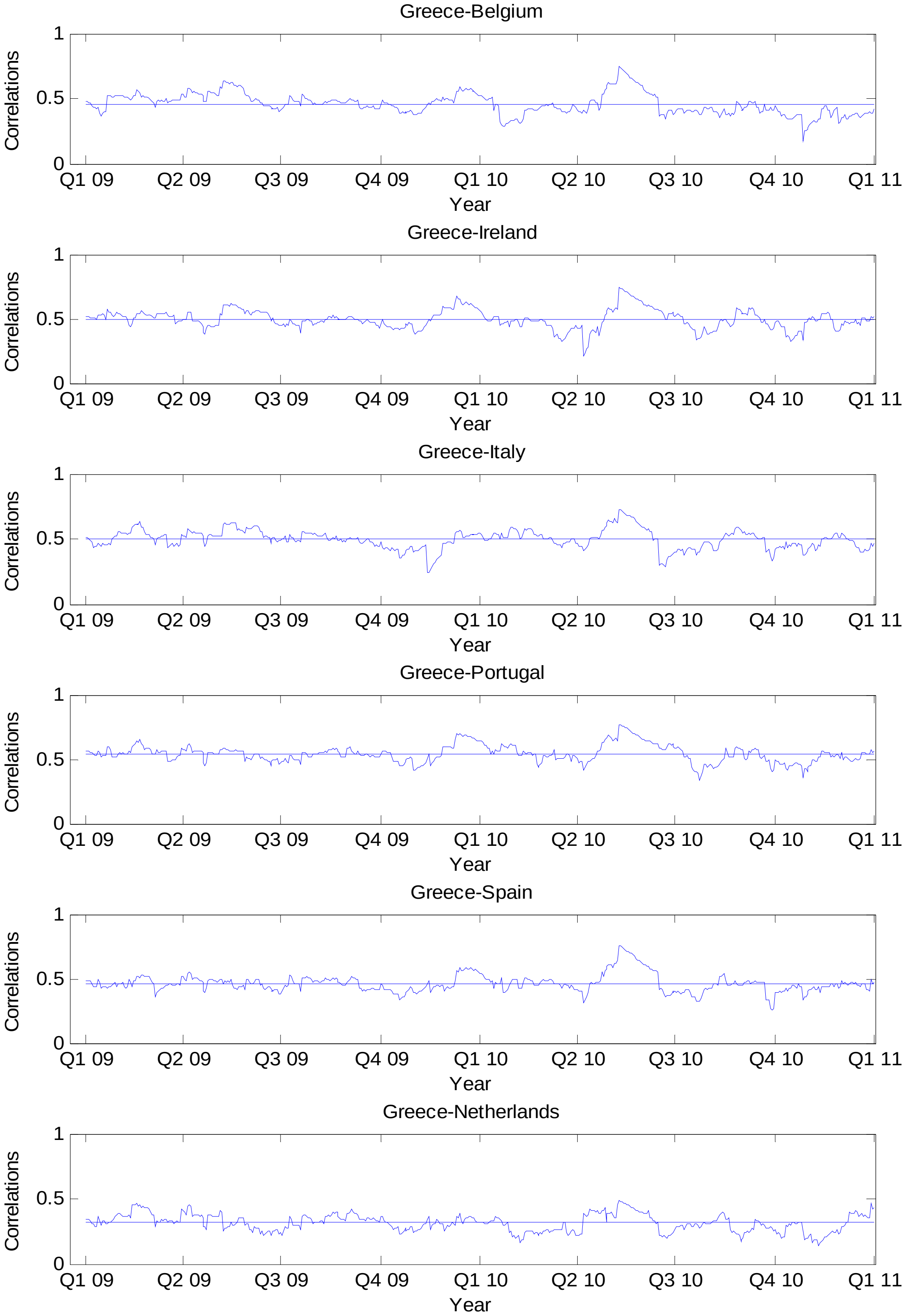

Figure 2: Greek correlation dynamics. Dynamic Conditional Correlations between yield spreads (to the German bond) of the relevant country pair.

4.2 Determining contagious effects

The analysis aims at investigating if refinancing problems of some European countries are due to contagious effects. If that was the case, some countries would suffer unjustified financial problems which were solely driven from deteriorated investor sentiment stemming from independent and bad news of other countries. As the sovereign debt crisis initially hit Greece, we take Greece as the origin of the crisis and examine if other countries suffer directly from the fact that Greece was in financial distress, even though they might actually be unrelated to the Greek problems and are in fact financially sound.

As an example consider Portugal. If it can be shown that contagion led to rising Portuguese government bond yields, then bad news about the Greek economic performance, competitiveness or indebtedness is extended to Portugal even though a higher Portuguese risk premium should not be fundamentally justified. Instead if no contagion is found, then the increase in Portugal's risk premium is economically and fundamentally justified and is not only caused by bad investor sentiment and panic introduced by the bad news about Greece.

According to much of the contagion literature, for the identification of contagion a strong increase in volatility adjusted cross-country correlation coefficients needs to be observed. As argued by Forbes and Rigobon (2002), a permanent increase in correlations which remain stable at the higher level once the increase is completed is not contagious but driven by stronger economic interdependences. Such economic integration is a time consuming process and does not revert back immediately. Consequently, contagious effects are identified only if correlation measures increase significantly during the contagious period, but do not remain permanently on that higher level.

The pairwise correlation dynamics show strong increases in the summer of 2010 for Belgium, Ireland, Italy, Portugal and Spain, giving rise to the assumption of potential mispricing of risk. Evidence for mispricing of sovereign default risk in Greece, Ireland, Italy, Portugal and Spain can also be found in Aizenman et al. (2011).
While the correlation coefficients fairly regularly bounce beyond and above the assumed constant correlation before Q2 2010 and after Q3 2010, the time period in between is characterized by a high increase of comovements. For all countries, the maximum of the correlations fall within that period. Also this increase is not permanent, as it reverts to the

assumed constant correlation clearly too fast as to argue for an economically driven increase. Consequently we argue that contagion effects can be identified in the Euro Area. Potentially existing fundamental problems were further worsened to at least some extent by fundamentally unrelated bad news about the Greek economy.

All correlation series in figure 2 display rather erratic behaviour. The important feature of the chart is the prominent spike in the summer of 2010. This spike showing considerably higher correlations occurs in the panels for the Greek yields to the Belgian, Irish, Italian, Portuguese and Spanish yields – but not in the panels for the correlations with the Dutch yields. Therefore we argue that contagion effects can be found from bad news about Greece to otherwise sound but maybe somewhat problematic countries, but are less likely or not existing to hit economically and politically rather stable countries. If however countries are under close investors' watch for various reasons, the sudden downturn in financing conditions of one observed country can cause spillover effects – or contagion – exaggerating the actual fundamental problems.

Summarizing, it can be concluded that the spreading refinancing problems of some European countries have to some extent been worsened by contagion and are not only based on suddenly deteriorating news about the competitiveness and fiscal stance of the countries in trouble. This conclusion is crucially important for the choice of politicy intervention. As argued by Forbes and Rigobon (2001) identified contagion effects infecting countries with no economically justified financing problems would infact call for some form of bail-out mechanism. Thereby, investors could be calmed down and refinancing costs possibly decrease to normal or fundamental values. This would allow the normal economic development of the country to continue without any detrimental effects from the contagion. Consequently, the bail-out capital would not be sacrificed in such a scenario as the stance of the borrowing economy is robust enough to allow for full and quick repayment. If, however, no contagion is identified, then the financing problems are mainly due to fundamental economic and fiscal problems of the relevant country. In such a situation a bail-out might only calm the investors down for a moment, but soon the economic grievance will reappear. The resulting renewed accentuation of financial distress would call for an additional bail-out, which however would again be useless for solving the fundamental problems of the country. Consequently, if there is no contagion at work a bail-out is unlikely to be successful and measures aiming at strengthening the competitiveness and structural reforms of the public debt and deficit levels of the country are presumably preferable.

For the current European situation this means specifically that rescue strategies should be adjusted to these insights. The approval or non-approval of a stabilising mechanism should – amongst other considerations – be certainly related to the identification and possibility of contagion effects. In May 2010 the European Financial Stability Facility was implemented and a 110 billion Euro loan to Greece was provided by the countries of the Eurozone and the IMF. As our results in figure 2 shows this was at a time in which a DCC-model identifies contagious effects at work and thus this decision seems indeed very reasonable. Further bail-outs should be evaluated with respect to the same or similar quantitative analysis.

4.3 Robustness

In order to check the robustness of the results observed so far, a similar analysis has been conducted using modified datasets. The DCC estimation of the correlation dynamics is also performed using the 10-year benchmark government bid yields instead of the bond yield spreads. Additionally, the 10 year CDS spreads between the seven analysed countries and Germany were implemented. All data is again used on a five-day week basis between 01/01/2009 and 12/31/2010 and is provided on Datastream.

The results of the robustness analysis are presented in Figure 3. The solid line represents the correlation dynamics for the bond yield spreads, the dashed line for the bond yields and the dotted line for the CDS spreads. The DCC models are specified to fit the new datasets.

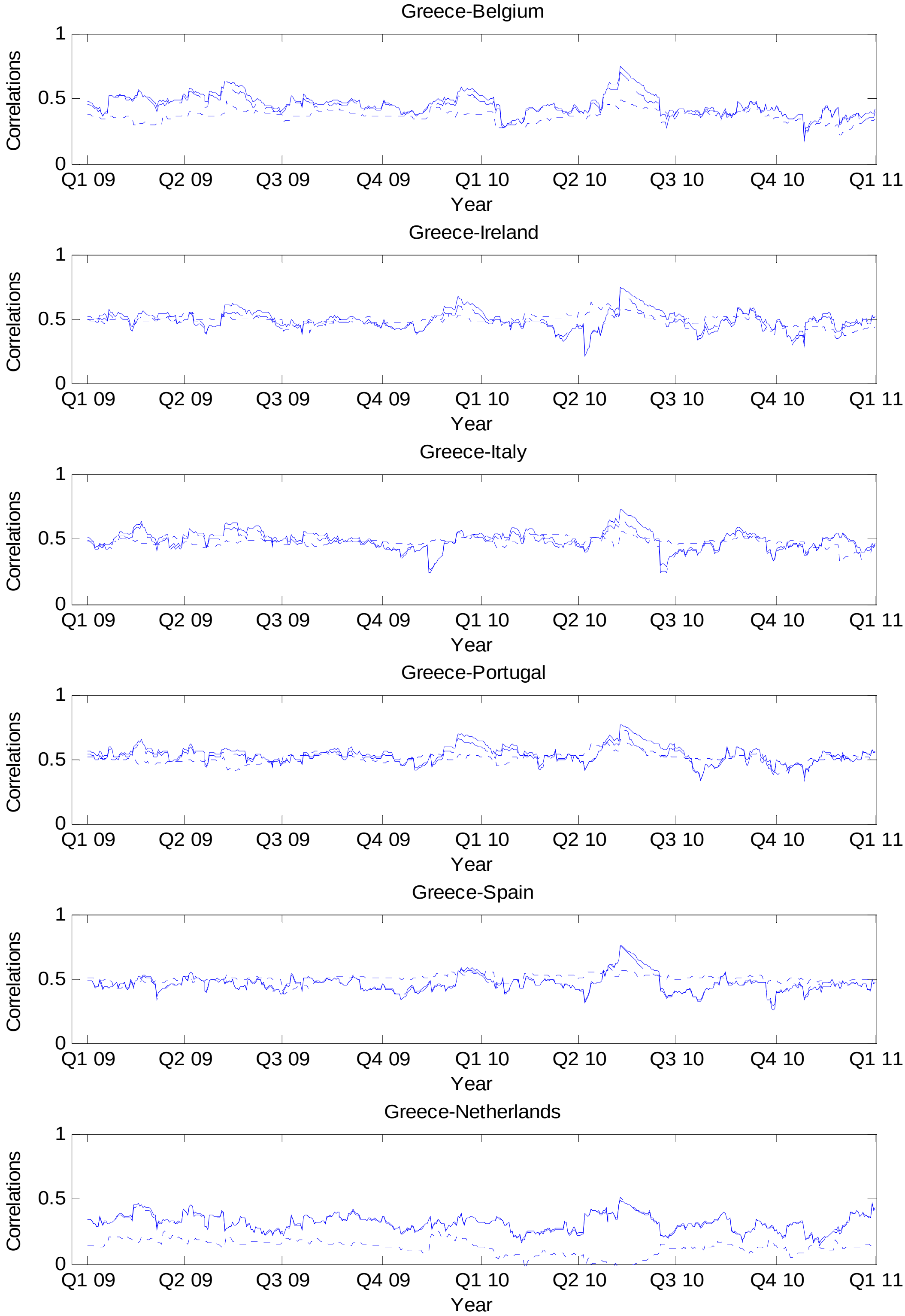

Figure 3: Greek correlation dynamics, robustness check. Solid line: Yield spread correlations. Dashed line: Yield correlations. Dotted line: CDS spread correlations.

It can be observed that the pattern of correlation dynamics remains roughly identical for all three time series. While a similar development with correlations peaking in summer 2010 can be observed for Belgium, Ireland, Italy and Portugal, the similarity of the dynamics is somewhat weaker for Spain. Most importantly the significant increase in correlations between Q2 2010 and Q3 2010 can be confirmed by using different datasets with different specifications. Consequently the results established in Subsections 4.1 and 4.2 seem to be quite robust. The time series of the Netherlands confirm the absence of contagion, as especially the correlation dynamics based on the CDS spread doesn't show its peak, but its lowest point in summer 2010.

5. Announcement Effects

So far we have shown that there seem to be contagious effects at work during the Euro Crisis in general. We now study if single Rating Agency announcements can by themselves trigger contagious effects. If a negative rating announcement in one country significantly increases cross-country correlations, this rating cut would then also influence the investors' sentiment about other countries in which there was no rating downgrade at all and in which – ceteris paribus – the fundamental data remained unchanged. In the following we investigate if negative rating announcements for Greece significantly changed the correlation dynamics and consequently altered the financial situation of the other countries analysed.

5.1 Model Setup

In order to analyse the contagion effects of announcements, univariate time series models for the DCC correlations are estimated and extended by including a rating announcement dummy. Taking Greece as the origin of the crisis, the correlation series between Greece on the one side and Belgium, Ireland, Italy, Portugal and Spain on the other side is implemented into an ARMA model. The country selection is due to the fact that those are identified to be affected by contagious effects in general and therefore it is interesting if announcements pronounce this phenomenon. A dummy variable indicating the negative rating announcements for Greece and a set of control variables according to (9) is introduced into the estimation equation.

$$\rho_t = \varphi + \Sigma_p \kappa_p \rho_{t-p} + \eta D_t + \Sigma_i \tau_i C_{i,t} + \Sigma_q \theta_q u_{t-(q-1)} \qquad (9)$$

In (9), ρ_t represents the dynamic correlation estimated in the DCC-equation (4), u_t the current innovations, D_t the rating announcement dummy for Greece, $C_{i,t}$ the set of the i = 1,…,I control variables and φ, the κ's, the θ's, τ's and η the parameters to be estimated.

5.2 Dataset

A dummy variable constructed with rating announcements for Greek sovereign debt between the period of 01/01/2009 and 12/31/2010 is used in order to test the impact of rating downgrades on correlations. During that time period only negative rating cuts were published. The dummy variable takes a value of one on each day on which Fitch, Moody's or Standard and Poor's announced a downgrade and a value of zero otherwise. For the whole sample there are 18 negative rating announcements.
The control variables are constructed with Fitch's sovereign debt ratings of the specific other country being analysed and discussed in more detail in the next section.[4]

5.3 Specification

In order to estimate equation (9) for the five correlation series the suitable ARMA-specification again needs to be identified. All time series are stationary and model selection for the levels is again conducted with Hannan-Rissanen model selection and Schwarz information criterion, models are checked using Portmanteau and LM tests. According to this procedure all correlation series follow an AR(1) process. The filtered correlations exhibit no sign for remaining autocorrelation or conditional heteroskedasticity.

Equation (9) is estimated for three different specifications. In the baseline scenario, only the Greek ratings dummy is included into the AR(1) models in order to test if such a rating announcement significantly influences the correlation dynamics. If a rating cut for Greece significantly increases the yield spread correlation between Greece and another country, one might conclude in favour of contagious effects. A country which is unrelated to Greece gets negatively affected by Greek rating deteriorations.

However, it might also be the case that the other country's rating is not independent from the Greek rating downgrade. For instance, if financial markets believe it has become more likely

[4] All publication dates can be directly obtained from the rating agencies' web sites and are available from the authors upon request.

that Portugal will be downgraded too, after a Greek downgrade, then investors would expect a subsequent Portuguese downgrade following the Greek rating downgrade. Hence, in this case a correlation increase between Greek and Portuguese yields would not be due to contagion or irrational investor sentiment, but by the rational investors' anticipation of an increased likelihood of a Portuguese rating cut. The worsened refinancing conditions of Portugal then do not result from announcement contagion, but from fundamental factors. Therefore, the second and third specifications try to control for the interdependence between Greek and other countries' ratings.

For the second specification a rating spread between the Greek and the five other countries' rating is used as control variable. Each rating is indexed to a number according to Afonso et al. (2011). Highest quality ratings (AAA ratings) receive a number of 17, very high credit risk and worse ratings (CCC+ and worse ratings) receive a number of 1, and all other ratings in between are linearly transformed to the number 2 – 16 accordingly. The rating spread is calculated by subtracting the Greek index from the different other countries' index. For the whole sample period, the Greek index is always smaller than other indices and therefore the control variable is positive. The smaller the spread turns out to be, the closer is the Greek rating to the compared rating. If it is more likely for similarly bad rated countries to obtain a rating cut once Greek was downgraded, then for such countries the control variable should have a positive impact on the correlation coefficients. Interdependences between the rating developments of two countries should hence be captured.

In the third specification, rating interdependences between two countries are captured by estimated dynamic correlations between those countries. Therefore, the rating development is again indexed and a DCC model is estimated for the ratings. In order to prepare the indexed ratings as suitable mean zero input variables for a DCC model, the rating time series are demeaned. Subsequently the simplest possible DCC specification with GARCH(1,1) and DCC(1,1) lag length selection is estimated. The resulting dynamic conditional correlations for the ratings are used as control variables accounting for the interdependence of rating developments. This third specification is however only feasible for Greece, Ireland, Portugal and Spain, as those are the only countries for which rating changes occurred between 01/01/2009 and 12/31/2011. Consequently, only rating correlation time series between Greece and Ireland, Greece and Portugal and Greece and Spain can be calculated, as correlation coefficients are not defined if one of the two variables of interest is constant.

5.4 Results

The conclusions of the contagion analysis of announcement effects are ambiguous. Equation (7) is calculated for specifications one and two for the correlations of Greece with Belgium, Ireland, Italy, Portugal and Spain, specification three for the correlations of Greece with Ireland, Portugal and Spain. The results are presented in Table 2.

Correlation Series	Parameter	1. Specification	2. Specification	3. Specification
Greece – Belgium	Rating Dummy	0.001	0.001	X
	Rating Spread	X	-0.001	X
	Rating Correlation	X	X	X
Greece – Ireland	Rating Dummy	0.010***	0.010	0.009
	Rating Spread	X	0.001	X
	Rating Correlation	X	X	-0.005**
Greece – Italy	Rating Dummy	0.001	0.001	X
	Rating Spread	X	0.000	X
	Rating Correlation	X	X	X
Greece – Portugal	Rating Dummy	0.007***	0.006**	0.007*
	Rating Spread	X	0.000	X
	Rating Correlation	X	X	-0.001
Greece – Spain	Rating Dummy	0.008	0.007*	0.007***
	Rating Spread	X	0.001	X
	Rating Correlation	X	X	-0.005***

Table 2: Greek announcement effect estimation: Dummy parameter (η) and control variable (τ) estimates. *, ** and * denote statistical significance at the 10%, 5% and 1% confidence level.**

As long as one assumes that the Greek ratings are independent from Irish or Portuguese ratings, the announcements of Greek rating cuts have a bad impact on Ireland and Portugal. The announcement dummy in specification one has a significantly positive effect on the correlation between Greece and the two countries. As the correlation of Greek and Irish or Portuguese bond spreads increases on Greek announcement days, the bad information about Greece spreads over to Ireland and Portugal and negative rating news on Greece seem to badly influence investors' perception of the financial stance of the two countries. Contagion can therefore be identified. If however one believes that the Greek and the Irish or the Portuguese rating are related to each other, it would be rational to expect a rating downgrade for Ireland and Portugal after Greece was downgraded. Therefore, contagion can only be identified if one controls for this increased downgrade probability. The hypothesis of contagion through rating downgrades is rejected for Ireland in specification 2 and 3, but accepted for Portugal in both specifications. In summary, the evidence for announcement contagion is quite clear for Portugal, but unclear for Ireland, however slightly favouring the existence of such effects.

For Spain no contagion can be shown in the baseline regression, however the dummy coefficients are significant both in specification two and three. Consequently, contagious effects are identified if the Greek and Spanish ratings are dependent on each other, otherwise not. Finally, we do not find significant announcement effects for Italy and Belgium.

Summarizing the analysis of Greek announcements we conclude that bad rating news show at least some tendency towards a generation of contagious effects for some countries. This tendency for correlation increases on announcement days is shown graphically for the Portuguese case in Figure 4.

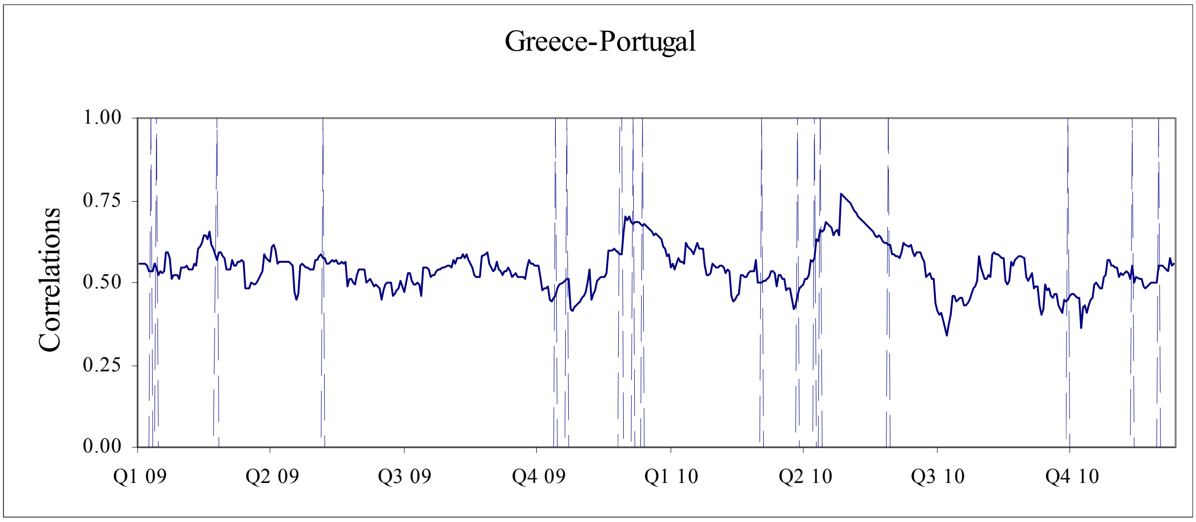

Figure 4: Announcement effect of Greek ratings. Solid line: Yield spread correlation between Greece and Portugal. Dashed line: Rating announcements for Greece.

The graphic shows the correlation dynamics between Greek and Portuguese yield spreads and indicates each day of rating announcements for Greece. For most of the announcement days it can be seen that the correlation tends to strongly increase with rating news.

The identification of contagious effects generated by rating announcements is important for different reasons. First, the rating development of different related countries needs to be kept in mind when it comes to interpreting bond yield movements or implementing measures aiming at influencing the bond markets. For instance countries which are badly affected by other countries' ratings should try to avoid the emission of new treasury bonds soon after downgrades of related countries as such news will put upward pressure on the required yield of their own new issue. Second, announcement effects are important from an investor's point of view (see e.g. Christiansen (2000)). The intraday behaviour of co-movements of different assets is important when it comes to risk management, asset allocation and asset pricing.

6. Conclusion

We have estimated a dynamic conditional correlation model (DCC) in order to analyse the correlation structure of Greek, Belgian, Irish, Italian, Portuguese, Spanish and Dutch bond yield spreads over the German yield to study contagion in the Euro Area. Our results do indicate the presence of contagious effects during the Euro Crisis. In particular, Belgian, Irish, Italian, Portuguese and Spanish yield spreads do increase along with their Greek counterpart. Thus it seems likely that Greek financial problems can spread via contagion to other Euro Area countries.

The resulting policy implications are ambiguous and should be drawn very carefully. While a bail-out, as implemented in summer 2010, can be regarded as a reasonable reaction to contagious pressures, the general development of bond markets of those countries also call for measures aiming at increasing their fiscal stance and competitiveness as well.

Finally, we studied if Greek rating downgrades generate contagious effects on other countries. We find that bad news about Greek ratings can in fact generate contagion to some other countries. However this does not hold for all countries in our sample as some are unaffected by Greek downgrades.

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